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Fuel Price Volatility and Asymmetric Transmission of Crude Oil Price Changes to Fuel Prices

Iuliana ZLATCU
Bucharest University of Economic Studies, Romania
iuliana.zlatcu@competitions.ro
Matei KUBINSCHI
Bucharest University of Economic Studies, Romania
matei.kubinschi@bnro.ro
Dinu BARNEA
Bucharest University of Economic Studies, Romania
dinu.barnea@bnro.ro

Abstract. *In recent years, retail fuel prices have raised the question of whether they are anomalous in relation with the crude oil prices. This paper aims to study the price volatility and the response of retail fuel prices to changes in international crude oil prices on the Romanian fuel market and on four other markets, namely Germany, France, Poland and Czech Rep., during the period between 2008 and 2014. By estimating univariate and multivariate GARCH models, as well as applying an MTAR threshold cointegration approach, the study finds similar patterns across the aforementioned markets. Our results show that the conditional volatility reached its peak during the outbreak of the financial crisis, together with an abrupt fall in time-varying correlations, which offer support for the asymmetric transmission hypothesis. Finally, the cointegration analysis detects some signs of asymmetric adjustments, supporting the hypothesis of a faster adjustment above the threshold in most cases.*

Keywords: Price volatility, GARCH, BEKK, asymmetric adjustment, threshold cointegration, fuel market, collusion.

JEL Classification: C32, D43, Q41.

1. Introduction

Retail fuel prices in Romania have consistently been a subject of concern for the media, politicians, economists and the public. Asymmetric price transmission is a particularly relevant economic subject, as it implies a welfare redistribution process between producers and consumers, with significant political and social implications. Consumers often claim that gasoline and diesel pump prices respond asymmetrically to crude oil price changes and that, in relation to crude oil prices, retail fuel prices rise faster than they fall. Therefore, the fuel market is often suspected of the existence of anticompetitive behaviour, such as cartel agreements.

In Romania, the fuel market is highly concentrated, with more than 70% of the market being controlled by four players. Due to the fact that there are many market conditions which favour the existence of collusion between undertakings, such as the low number of players, the homogeneous nature of the products, inelastic demand, high entry barriers or vertical integration, competition problems may easily occur on this market.

In the past two years, the Romanian Competition Council (RCC) has made an assessment of some industries from the perspective of the Aggregate Index of Competition Pressure, an instrument that aims to identify markets that are prone to anticompetitive behaviour. According to the RCC analysis, the retail fuel market is characterized by a low degree of predisposition to competition. Moreover, the RCC uncovered a cartel on the gasoline retail market and sanctioned six undertakings in 2011 for being involved in an agreement aimed at stopping the sales of Eco Premium gasoline. The companies, which were in control of 90% of the market, received a fine of over 200 million euro. Finally, an investigation has been opened since 2005 on the fuel market concerning possible price fixing. Hence, we consider a further analysis in this sector to be important.

Price volatility and transmission have been widely investigated over time, in relation to the fuel market. Bolotova, Connor and Miller (2005) study the impact of the lysine and citric acid cartels on price level and variance. Using a GARCH model to analyse the price volatility patterns for lysine and citric acid cartels, their study finds that the lysine price increased by over 25% and its variance decreased during the cartel period; the citric acid conspiracy increased prices by 9 cents per pound relative to pre-cartel and post-cartel periods. In Esposito and Ferrero (2006), the authors test the variance screen on two markets, motor fuel market and personal care and baby food market, to see whether a price variance screen could have identified collusion in two Italian cartel cases. They compared price volatility across different markets in Europe and found that Italy's standard deviations were the lowest for both gasoline and diesel fuel prices. The authors concluded that the cartel detection method based on price variance would have successfully detected collusion prior to the findings of the Italian Competition Authority.

Furthermore, the relationship between the price of crude oil and the pump price has been extensively studied in recent literature. Several methods have been used to study the transmission between crude oil price and pump prices. Grasso and Manera (2007) study price asymmetries for five European countries, using three different methods, namely asymmetric error-correction models (ECM), autoregressive threshold ECM and ECM with threshold cointegration. All these models find a temporal delay in the reaction of retail prices to oil price changes, as well as some evidence of asymmetric behaviour. Regarding the fuel market in Germany, Asane-Otoo and Schneider (2014) also investigate the international crude oil prices to pump price transmission and, at national level, find positive asymmetric relationships for both gasoline and diesel for the 2003-2007 period, indicating that pump prices react faster to increases, while for the period 2009-2013 they detect negative asymmetry for retail prices. Bermingham and O'Brien (2010) use threshold autoregressive models and monthly data for the period 1997-2008 to examine the Irish and UK liquid fuel market and find no evidence for the asymmetric price hypothesis.

Regarding the asymmetric price adjustment, several explanations have been advanced (OECD, 2013): market power and tacit collusion, search costs, adjustment costs in refining and wholesale, and inventory management by consumers. Market power is one of the most important causes identified in the literature for the asymmetric price transmission (Balke et al., 1998). As the fuel market has an oligopolistic structure, the players use their market power in the pricing process; in consequence, when the international oil price increases, the players tend to fully transmit the increased cost to the fuel price and, when faced with declines in input costs, they have less incentive to reduce prices. Instead of competing with each other, players in an oligopolistic market may decide to collude, tacitly or explicitly. Of course, given the fact that the market in discussion is an oligopoly, the asymmetric price transmission can be the result of a natural parallel behaviour, an agreement being unnecessary, in this case.

In the present paper, we study price volatility on the retail fuel market and investigate the relationship between retail prices and crude oil prices. This is done in order to assess whether a change in crude oil will be transmitted to the pump fuel prices on the Romanian market, as well as on the other markets included in this study, namely Germany, France, Czech Republic and Poland, over the period between 2008 and 2014.

2. Methodology

In this section we discuss the econometric methodology used to study price volatility and the existence of asymmetric pass-through between crude oil prices and fuel pump prices. Specifically, we discuss the Generalized Autoregressive Conditional Heteroscedasticity (GARCH) model, both in the univariate and the multivariate versions, and the threshold autoregressive (TAR) cointegration model.

3.1. Modelling volatility

3.1.1. Generalized Autoregressive Conditional Heteroscedasticity (GARCH) model

In order to study price volatility, we employ an econometric model that has the advantage of capturing some stylized facts of the time series, such as time-varying volatility, namely the generalized autoregressive conditional heteroscedasticity (GARCH) model, introduced by Bollerslev (1986). The econometric formulation allows for the conditional variance of u at time t to depend not only on the squared error term in the previous time period, but also on its conditional variance in the previous time period.

The conditional variance of a GARCH (q,p) that depends on q lags of the squared error and p lags of the conditional variance is as follows:

$$\sigma_t^2 = \alpha_0 + \sum_{i=1}^{q} \alpha_i u_{t-i}^2 + \sum_{j=1}^{p} \beta_j \sigma_{t-j}^2 \tag{1}$$

For this model to be well defined and the conditional variance to be positive, α_1,..., α_q, β_1,...., β_p, and α_0 must be positive constant parameters.

Interpreted in a financial context, this model describes how an agent tries to forecast volatility for the next period based on long-term average (α_0) variance, information on the volatility observed in the previous period (ARCH term) and the previous period variance value (GARCH term).

The simplest form of the conditional variance equation is the GARCH (1, 1) model, which can be written as:

$$\sigma_t^2 = \alpha_0 + \alpha_1 u_{t-1}^2 + \beta_1 \sigma_{t-1}^2 \tag{2}$$

It is widely considered that a GARCH (1, 1) specification is sufficiently robust to capture the evolution of volatility, although many extensions and formulations of the model exist in related econometric literature.

3.1.2 Baba-Engle- Kraft-Kroner (BEKK) model

Obtaining estimates for the dynamic correlations established between the analysed time series implies applying a multivariate GARCH approach, namely the Baba-Engle-Kraft-Kroner (BEKK) model, defined in Engle and Kroner (1995). The BEKK model is designed as a multidimensional extension of the univariate GARCH model and has the desirable property that the conditional covariance matrices are positive definite. The BEKK (q,p,k) representation is the following:

$$y_t = \mu_t + \varepsilon_t, \tag{3}$$

$$\varepsilon_t = H_t^{1/2}\eta_t, \quad \eta_t \sim iid(0, I_n) \tag{4}$$

$$H_t = CC' + \sum_{j=1}^{q}\sum_{k=1}^{K} A'_{kj}\, r_{t-j} r'_{t-j} A_{kj} + \sum_{j=1}^{p}\sum_{k=1}^{K} B'_{kj}\, H_{t-j} B_{kj} \tag{5}$$

where A_{kj}, B_{kj} and C are *N x N* parameter matrices, C is lower triangular to ensure the positive definiteness of H_t and the summation limit K accounts for the generality of the autoregressive process. The decomposition of the constant term into a product of two triangular matrices is required to ensure positive definiteness of H_t. Engle and Kroner

(1995) found that the BEKK model is covariance stationary if and only if the eigenvalues denoted by $\sum_{j=1}^{q}\sum_{k=1}^{K} A_{kj} \otimes Akj + \sum_{i=1}^{p} B_{kj} \otimes B_{kj}$ are less than one in absolute values, where $\otimes$ denotes the Kronecker product of two matrices. The computational requirements for the estimation procedure of a BEKK model are relatively high due to the fact that the process involves estimating $(p+q)KN^2 + N(N+1)/2$ parameters, therefore being normally assumed that $p = q = k = 1$ in numerical applications of the model. For instance, in the simplest case of a BEKK (1,1,1) model, there are 11 parameters that need to be estimated. Model estimation for all classes of multivariate GARCH models is performed using maximum likelihood with the following log-likelihood function:

$$\mathcal{L}(\theta) = -\frac{TN}{2}\log(2\pi) - \frac{1}{2}\sum_{t=1}^{T}(\log|H_t| + r_t' H_t^{-1} r_t) \tag{6}$$

where θ is the vector that contains all the parameters, N is the number of variables and T is the number of observations in the time series. The time-varying correlations between each of the analysed time series can be derived by dividing the conditional covariances by the product of conditional standard deviations obtained from the BEKK model. The numerical computation and optimization procedures were done in Matlab 2013a using the Oxford MFE toolbox written by Kevin Sheppard.

3.2. Modelling Asymmetry

3.2.1. Threshold autoregressive (TAR) model

Our study uses the threshold cointegration approach formulated by Enders and Siklos (2001), to investigate whether asymmetric transmission exists between the crude oil prices and the fuel retail prices. The Enders and Siklos test extends the Engle and Granger (1987) approach to test for a long-run equilibrium relationship by allowing for asymmetric price adjustments. The test focuses on the residuals from the estimation of the long-run causal relationship between the fuel retail price and crude oil price, both integrated of order one.

$$x_t = \delta + \beta y_t + \mu_t \tag{7}$$

$$\mu_t = x_t - \delta - \beta y_t \tag{8}$$

where x_t is the fuel retail price, y_t is the crude oil price and μ_t is the residual that captures the deviation from the long-run equilibrium. For the series to be cointegrated, μ_t needs to be stationary. The Augmented Dickey-Fuller (ADF) test can be used to ascertain whether the residuals are stationary and the series are cointegrated. In other words, if $-2<\rho<0$ in the equation:

$$\Delta\mu_t = \rho\mu_{t-1} + v_t \tag{9}$$

then the two variables admit an error-correction representation in the form:

$$\Delta x_t = \alpha(\delta - \beta y_{t-1}) + \varepsilon_t \tag{10}$$

Following Enders and Siklos (2001), to allow for asymmetric adjustments, the residual μ_t is modeled as a threshold autoregressive (TAR) process that can be written as:

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + v_t \tag{11}$$

where τ is the value of the threshold and I_t is the Heaviside indicator function that depends on the lagged values of the residual μ_{t-1}:

$$I_t = \begin{cases} 1 \ , \ if \ \ \mu_{t-1} \geq \tau \\ 0 \ , if \ \ \mu_{t-1} < \tau \end{cases} \tag{12a}$$

The same paper suggest another alternative, that the adjustment process be dependent on the change in μ_{t-1} in the previous period rather than the state of μ_{t-1}. In this case, a momentum threshold autoregressive (M-TAR) model turns (12a) into:

$$I_t = \begin{cases} 1 \ , \ if \ \ \Delta\mu_{t-1} \geq \tau \\ 0 \ , if \ \ \Delta\mu_{t-1} < \tau \end{cases} \tag{12b}$$

The only difference between the TAR specification and the M-TAR specification is the definition of the Heaviside indicator function (I_t); the Heaviside indicator function is based on the level value of the threshold indicator variable in the former and on the change in the threshold indicator variable in the latter.

3. Data and results

For the purpose of the present study we use three datasets spanning from January 2008 to December 2014. First, for the retail prices, we use weekly prices for gasoline and diesel, obtained from the European Commission. The prices are net of taxes and are expressed in Euro per 1000 litres; secondly, for the crude oil prices, we use weekly Europe Brent spot price (FOB) from the Energy Information Administration (EIA) of the U.S. Department of Energy. Prices are expressed in US dollar per barrel. In order to construct a data set comparable with the retail prices, we convert them to Euro per 1000 litres using the appropriate exchange rate and correspondence (1 barrel =158.987295 litres).

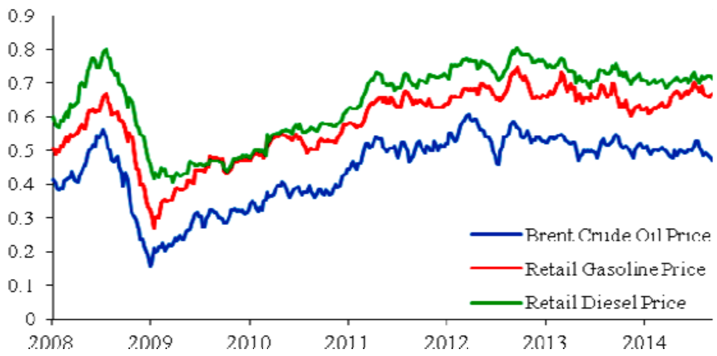
Our study focuses on two fuel products: 95 octane gasoline and diesel oil. These products are widely bought by the consumers and are basically homogeneous products.

The price of crude oil in international markets is considered the most important factor in the gasoline and diesel prices because it represents the largest component of the producing and marketing cost. World crude oil prices are established in relation to three market traded benchmarks (OECD, 2014): West Texas Intermediate, Brent Blend and Dubai/Oman. The Brent crude oil is extracted from the North Sea and is the most imported by the European countries.

In Romania, in 2014, crude oil price accounted for about 73% of gasoline average retail price (excl. taxes), and 67% in the case of diesel average retail price (excl. taxes), the rest being represented by refining, distribution and marketing costs.

Figure 1 shows the relationship between the price of gasoline and diesel on the one hand and the price of the crude oil on the other, over the last seven years, illustrating that the series move closely together. However, retail prices seem less responsive to changes in crude oil in certain periods.

Figure 1. *Weekly retail fuel and crude oil prices (Euro/litre)*



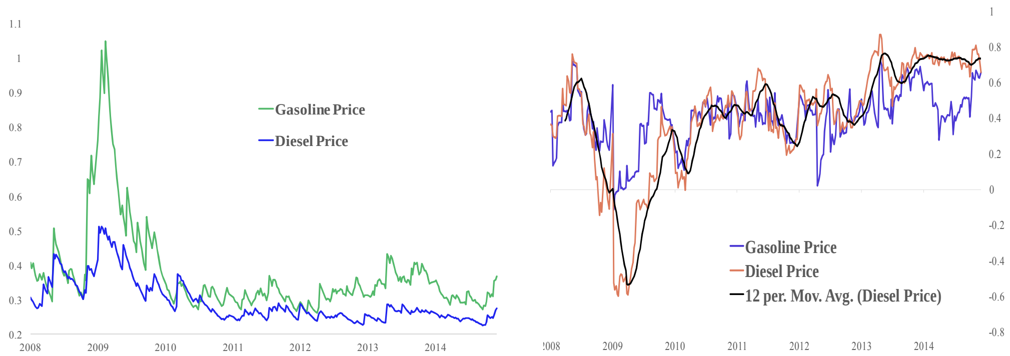
The study includes Romania and four other countries from the European Union: two core economies of the EU (Germany and France) and two other Central and Eastern European countries (Czech Rep. and Poland) as they share similar economic backgrounds with Romania.

An initial analysis of the data is made to assess whether the series are stationary. Accordingly, Augmented Dickey Fuller (ADF) and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) stationarity tests are conducted. As the price series are not stationary, weekly percentage changes were used. Secondly, we t*est* for *cointegration using* the Johansen methodology to validate the existence of a long term relationship between variables. The test results show that crude oil and retail price series are cointegrated.

In what follows, we study the price volatility for gasoline and diesel fuels. Results from estimating the volatility using GARCH are reported in Figure 2.

We observe that price volatility displays similar patterns across retail gasoline and diesel prices over the considered timeframe. Both price volatility indices encounter a steep rise at the beginning of 2008, with different magnitudes of growth. While diesel price estimated conditional volatility records a maximum value of 50%, during the outbreak of the global financial crisis, the gasoline volatility index exhibits a sharp rise, reaching over 100% annualized daily volatility, during the same turmoil episode.

Figure 2. *Gasoline and diesel retail prices estimated annualized volatility (left chart) and time-varying BEKK correlations (right chart) for the Romanian market*



To go further and improve the understanding of fuel price volatility, we study to what degree the retail fuel prices are related to the crude oil prices. Consequently, we estimate BEKK models for gasoline, diesel and oil prices in order to capture the dynamic time-varying correlations. The main finding is that conditional correlations exhibit significant changes over time for both gasoline and diesel. Firstly, analysing the trend displayed over the 2008-2014 period reveals an upward tendency for the majority of countries included in the sample, leading to the conclusion that the deepening of financial markets, which is a natural consequence of the globalization process, has led to a more efficient transmission of international oil prices.

Secondly, focusing on the time-varying characteristics displayed by the conditional correlations, we can observe a similar pattern emerging across the entire sample during international turmoil episodes, such as the global financial crisis of 2008 or the European sovereign debt crisis. The significant oil price market corrections are accompanied by a sudden drop in conditional correlation, suggesting a lagged effect on retail diesel and gasoline prices. Naturally, it is clear that commodity market prices are much more sensitive to global shocks and that adjustments in retail prices are not instantaneous and depend on several other firm and economy related factors. Nonetheless, these temporary slumps in conditional correlations highlight the complexity of the oil-retail price relationship and, subsequently, can offer support in favour of the asymmetric transmission hypothesis. In the absence of external shocks, retail price dynamics should be driven by maximizing utility functions in the short run (via rational economic agents pricing mechanism) and by the resource scarcity, in the long run.

Further, we compare price behaviour across different markets in Europe. Results are presented in Figure 3 (for estimated volatility) and Figure 4 (for time-varying correlations).

Figure 3. *Gasoline and diesel retail prices estimated annualized GARCH volatilities for selected European market*

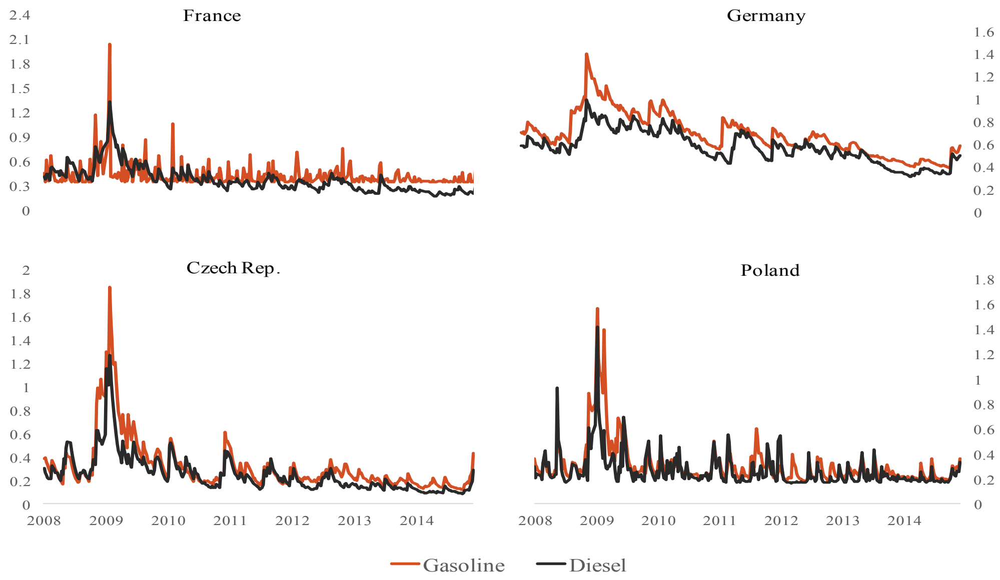
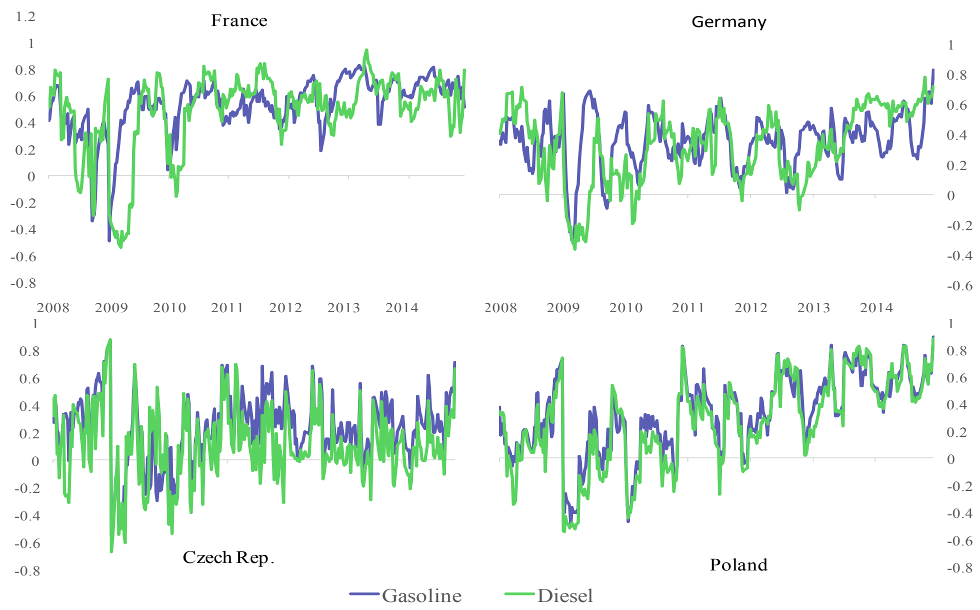


Figure 4. *Estimated time-varying BEKK correlations for selected European markets*



To capture possible asymmetric adjustments of retail fuel prices in response to crude oil prices, we use the M-TAR cointegration model. The results from the threshold cointegration analysis are reported in Table 1.

Table 1. *M-TAR estimation results*

	ρ_1	ρ_2	t-max	Φ-stat	F-equal
Romania gasoline	-0.128	-0.077	-2.214	7.932	1.078
			(-1.9958)	(6.2516)	(3.6641)
Romania diesel	-0.141	-0.114	-2.747	9.35	0.235
			(-1.9943)	(6.2384)	(3.7578)
Germany gasoline	-0.305	-0.154	-2.556	14.086	3.967
			(-1.985)	(6.1592)	(3.9088)
Germany diesel	-0.143	-0.108	-2.274	6.141	0.306
			(-1.9578)	(6.2119)	(3.7345)
France gasoline	-0.237	-0.296	-3.718	13.923	0.547
			(-1.9781)	(6.3044)	(3.7171)
France diesel	-0.092	-0.015	-0.442	4.185	2.601
			(-1.9796)	(6.1991)	(3.9167)
Czech gasoline	-0.088	-0.15	-2.292	9.615	1.427
			(-1.9706)	(6.2729)	(4.128)
Czech diesel	-0.047	-0.048	-1.52	2.292	0.001
			(-1.940)	(6.1476)	(3.6420)
Poland gasoline	-0.111	-0.077	-2.355	7.379	0.573
			(-1.9556)	(6.2099)	(3.8045)
Poland diesel	-0.174	-0.135	-3.098	12.328	0.471
			(-1.9948)	(6.3722)	(3.8053)

For our estimation, the threshold has been set to 0, while the critical values for each test have been generated through 10,000 Monte-Carlo simulations. Two t-statistics have been determined for $\rho_1=0$, and $\rho_2=0$ respectively, and one F-statistic, called Φ for the joint hypothesis of $\rho_1=\rho_2=0$. *T-max* denotes the largest of the t-statistics and, since both coefficients need to be negative in order for convergence to be achieved, *t-max* is an indicator of both conditions. However, Enders and Siklos (2001) find that the *Φ-statistic* can have substantially more power than the *t-max* statistic.

The critical values reported for the *t-max* statistic at the 95% level vary from -1.9 and -2 in all cases. The statistic value exceeds these levels only in the case of French diesel and Czech diesel, meaning that for all other series, the null hypothesis of no cointegration can be rejected at the 5% significance level.

Similarly, the results of the Φ test are very close, indicating one other missing cointegration relationship, in the case of Germany diesel, despite the statistic value being very close to the 95% critical value.

Enders and Siklos find however that, if the adjustment is nearly symmetric ($\rho_1 \approx \rho_2$), the Engle-Granger method for testing for cointegration has more power than the Φ and *t-max* statistics. In the case of asymmetric error corrections, the *Φ-statistic* should be used over the other two, with the *t-max* test giving the least reliable results in all cases.

In the cases where cointegration could not be rejected, the next step is testing for asymmetry in the adjustment relationships; in other words, we test whether an asymmetric adjustment model describes the relationship between the two variables better than a regular error-correction model. The results of the F-test for the null hypothesis $\rho_1=\rho_2$ are presented under the F-equal column. The null can be rejected at a 95%

confidence level only in the case of German gasoline, where negative discrepancies seem to persist longer, as the coefficient ρ_1 of the above threshold adjustment is higher in absolute value. For Romania, the adjustment is almost symmetric in the case of diesel, while having a slight asymmetry for gasoline, indicating that gasoline prices adjust faster for positive deviation from equilibrium than for negative deviations.

4. Conclusions

The present paper studies fuel price volatility and the transmission mechanism of crude oil price changes to retail gasoline and diesel prices focusing on five European markets, namely Romania, Germany, France, Poland and Czech Rep. during the period between 2008 and 2014. By estimating univariate and multivariate GARCH models, as well as an MTAR threshold cointegration approach, we find similar patterns across the aforementioned markets. Our results show that the conditional volatility reached its peak during the outbreak of the financial crisis, together with an abrupt fall in time-varying correlations which offer support for the asymmetric transmission hypothesis. The cointegration analysis detects some signs of asymmetric adjustments, but these are mostly statistically insignificant, giving away clues of a faster adjustment above the threshold in most cases. Summing up, the similar behaviour combined with the moderate variation in the price volatility suggest an efficient functioning of the fuel market, despite the findings not ruling out the existence of collusion on isolated market segments in certain periods. Future developments on the subject of price asymmetry in the fuel industry should focus on expanding the present approach by taking into account additional factors such as industry segments and size, dataset quality and estimation techniques.

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